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Abstract

This paper exploits the significant recent variation in United States house prices to empirically examine the effect on housing equity constraints and nominal loss aversion on household mobility. The analysis uses unique, detailed data from 1985-1996 on household characteristics, mobility, and wealth from the National Longitudinal Survey of Youth (NLSY79) matched with house price data from 149 metropolitan areas to estimate semiparametric proportional hazard models of intra- and intermetropolitan mobility. There are five principal findings. First, household intrametropolitan own-to-own mobility responds differently to nominal housing losses than to gains. Second, nominal loss aversion is significantly less pronounced in intrametropolitan own-to-rent and intermetropolitan mobility, respectively. Third, there is some evidence of binding equity constraints in intrametropolitan own-to-own mobility. Fourth, there is little evidence that low equity constrains intrametropolitan own-to-rent and intermetropolitan mobility, respectively. Fifth, a comparison of the estimated effects indicates that nominal loss aversion has a more dominant effect than equity constraints in restricting household mobility, roughly two and one-half to three times the impact of equity constraints.

I. Introduction

Housing markets often exhibit behavior that cannot be explained by standard asset-market models (Poterba 1984). For example, they display rapid swings in prices, strong positive correlation of prices and trading volume over the housing cycle, and the observed reluctance of prospective sellers to reduce asking prices in down markets (Stein 1995; Genesove and Mayer 1997, 2001). An important part of recent research has been to propose and empirically test new theories that explain these puzzles. Two seemingly related but competing theories have emerged. The first is housing equity (or collateral) constraints, analyzed by Stein (1995), Genesove and Mayer (1997), Henley (1998a), Lamont and Stein (1999), Ortalo-Magne and Rady (1998), and Chan (2001), among others. The second, and quite provocative, theory is nominal loss aversion, analyzed first by Genesove and Mayer (2001).

Although both theories rely on the same propagation—a decline in nominal house prices—they have distinctly different implications for housing market behavior and government policy. Equity constraints occur because of down payment requirements in mortgage lending. Specifically, because most home purchases are mortgage financed, housing is a highly leveraged asset. A nominal price decline can result in equity- (or down payment-) constrained households who cannot move, which decreases market demand and results in further price declines that further constrain household mobility. In addition, down payment requirements arise for a number of reasons (Engelhardt 1996b). First, households with equity in a home share with the lender the risk of a marketwide decline in house prices. Second, down payments reduce moral hazard in the maintenance of the home and its value. Third, down payments help curtail adverse selection from asymmetric information in the mortgage lending market. These reasons all point to some sort of market failure that, in principle, could be addressed through government policy to ease binding equity constraints. In contrast, nominal loss aversion, whereby households are averse to realizing nominal housing market losses and, hence, treat gains and losses asymmetrically, is a characteristic of preferences, which typically are not thought of as affected by government policy instruments.

This paper empirically examines the effect of equity constraints and nominal loss aversion on household mobility. The data used are on young homeowners from the National Longitudinal Survey of Youth (NLSY79). Young homeowners are an ideal group to study. They have high job and geographic mobility and are highly leveraged. They are the most susceptible to equity constraints to mobility. In addition, the analysis exploits the significant metropolitan variation in housing market performance in the United States for 1985-1996. This period encompasses the well-known recessions in the energy states, the Northeast, and California, as well as the rising economic tides in the South, Midwest, and Pacific Northwest. These data display rich variation in nominal losses and gains.

Moreover, the data are unique. Detailed data on demographics, employment, and wealth come from the public-use version of the NLSY79. These were matched to administrative address data to construct mobility histories that are more detailed than those available from the restricted-access NLSY79 Geocode data. In addition, housing value and mortgage data from the NLSY79 master file were used. These data are top-coded on the public-use version of the data. The NLSY79 data were matched with weighted-repeat-sales house price indices for 149 metropolitan areas provided by the Federal National Mortgage Association (Fannie Mae) and the Federal Home Loan Mortgage Corporation (Freddie Mac). Analysis with data drawn from a broad sample of metropolitan areas is an important contribution of this paper, because most of the existing empirical analyses of equity constraints and nominal loss aversion used data from the New York metropolitan area (Chan 2001) and the downtown Boston condominium market (Genesove and Mayer 1997, 2001), respectively.

The final data set was used to estimate semiparametric proportional hazard models of intra- and intermetropolitan mobility. There are five primary findings. First, household intrametropolitan own-to-own mobility responds differently to nominal housing losses than to gains. Evaluated at the sample mean, having experienced a nominal gain reduced the baseline hazard of such a move by 37 percent. That is, households hold on to housing market gains. In contrast, having experienced a nominal loss *reduced* the baseline hazard of such a move by 27 percent. That is, households hold on to housing market losses. In all specifications, gains are treated differently from losses. These effects were estimated controlling for equity constraints, so that these results may be interpreted as evidence of nominal loss aversion. These findings confirm those in Genesove and Mayer (2001), and complement the results of Engelhardt (1996a), who found evidence of the asymmetric treatment of nominal housing gains and losses on homeowner saving behavior. Second, nominal loss aversion is significantly less pronounced in intrametropolitan own-to-rent and intermetropolitan mobility, respectively. For these types of mobility, households hold on to nominal housing market gains, but losses have little economic and statistical effect on mobility. Third, there is some evidence of binding equity constraints in intrametropolitan own-to-own mobility (although weaker in a statistical sense than that for loss aversion). Evaluated at the sample mean, being a constrained household (with less than 20 percent equity) reduced the baseline hazard of such a move by 11 percent. This complements the findings of Chan (2001) on mobility, as well as Caplin, Freeman, and Tracy (1997), who argued persuasively that low housing equity constrained residential mortgage refinancing, reduced consumption, and exacerbated regional recessions. Fourth, there is little evidence that low equity constrains intrametropolitan own-to-rent and intermetropolitan mobility, respectively. Fifth, a comparison of the estimated effects indicates that nominal loss aversion has a more dominant effect than equity constraints in restricting household mobility, roughly two and one-half to three times the impact of equity constraints. This result complements the finding of Genesove and Mayer (2001) for the downtown Boston condominium market.

The results in the current paper likely have some broader implications. In a provocative series of papers, Oswald (1996, 1997, 1999) has argued that housing market impediments in general can explain the significant cross-country differences in unemployment rates in the OECD countries, cross-regional differences in Britain and Canada, and cross-state differences in the United States.[1] While it seems unlikely a priori that there is a single explanation for the unemployment problem, if Oswald's hypothesis is correct, the implications are far-reaching. Because most countries intervene in housing markets heavily through tax subsidies to home ownership, rent control, public housing, zoning, growth restrictions, and mortgage market regulation, government housing policies could have significant labor market repercussions. Although the current paper was not intended as a direct test of the Oswald hypothesis, two results cast doubt on its relevance for the United States. First, low housing equity only appears to constrain homeowner mobility within metropolitan areas, but not *between* metropolitan areas. Second, homeowners who are not employed are significantly more likely to move between metropolitan areas. These results are inconsistent with Oswald (1996, 1997, 1999), but are not inconsistent with the influential work of Blanchard and Katz (1992), who argued that the primary mechanism for economic adjustment across regions in the United States is through the migration of labor, not the adjustment of market wages.

The paper is organized as follows. Section II describes the economic relationship between housing equity, nominal loss aversion, and mobility, and discusses related empirical work from the literature. Section III discusses the data and preliminary analysis. Section IV outlines the hazard model specification. Section V discusses the estimation results. There is a brief conclusion.

II. Previous Literature

The basic theory of equity constraints was laid out in Stein (1995) and is illustrated most clearly with an example. Consider a household that purchases a $100,000 home with a 10 percent down payment.[2] If house prices rise by 10 percent, the home is worth $110,000 and the household has $20,000 in equity (or 18.2 percent). For the same down payment requirement, the household could trade up significantly: it now could use that equity to purchase a $200,000 home with 10 percent down. However, with a 10 percent decline in prices, the home is worth $90,000 and the household has no equity. The household could not make the down payment on the same home without other wealth.[3] With no other wealth, the household cannot move and remain a homeowner.[4] Thus, nominal house price changes can have asymmetric effects on mobility: households can lever capital gains to purchase larger homes, but they become constrained by capital losses.[5]

Stein (1995) formalized this intuition into a liquidity-based model of the housing market. He showed that this asymmetric effect can result in within-equilibrium housing market multipliers and multiple equilibria. These results hold even if constrained households have the option of moving and renting. The strength of the multipliers depended on the fraction of owners who were constrained movers—only in markets with a sufficiently large fraction of constrained owners could there be significant feedback effects to house prices.[6] Ortalo-Magne and Rady (1998) have generalized Stein's findings in an overlapping-generations framework.[7]

Recent research on Boston and New York has provided some evidence in support of equity constraints.[8] One implication of Stein's model is that constrained owners may "go fishing," i.e., offer their houses for sale at an above market price that would allow a move in the (low probability) event that a buyer arrives to pay that price. Genesove and Mayer (1997) examined the effect of equity on the time-to-sale and listing behavior of potential sellers in the downtown Boston condominium market. They found that constrained sellers (i.e., with less than 20 percent equity) were more likely than unconstrained sellers to ask above market prices, and that this resulted in an inverse correlation between prices and time-to-sale.

Chan (2001) has provided the only direct evidence on equity and mobility in the United States. She examined the experience of homeowners in the New York metropolitan area with a unique data set of Chemical Bank adjustable-rate mortgages. She found that constrained households (i.e., less than 20 percent equity) experienced a 24 percent reduction in mobility relative to unconstrained owners in the four years after the decline in prices. Her study is noteworthy in that the mortgage data were of unusually high quality, with arguably no measurement error in the mortgage spell length and explanatory variables. A "move" was defined as a mortgage termination. Chan argued that ARM refinancing was rare because of a low-cost option of conversion to a fixed-rate loan at the market rate, and, as such, mortgage duration was a good proxy for residence duration. In addition, she provided evidence from public records that 95 percent of ARM terminations in her sample were from moves.[9]

Nominal loss aversion is a central feature of the prospect theory of Kahneman and Tversky (1979). Barberis, Huang, and Santos (2001) have integrated nominal loss aversion into an asset pricing model that explains the large equity premium, volatility in expected returns, and level of average return in the stock market better than traditional consumption-based models. Genesove and Mayer (2001) were the first to examine nominal loss aversion in the housing market. The basic hypothesis is that homeowners treat gains and losses differently, and are reluctant to realize nominal losses; hence, they will set higher list prices and have longer time on the market in the hope that they will find a buyer with an offer high enough to attenuate the nominal loss. Genesove and Mayer (2001) used similar (but updated) data from the downtown Boston condominium market as in Genesove and Mayer (1997). However, in the updated analysis, they found that most of seller behavior seemed to be driven by nominal loss aversion.

Only about one-quarter of the effect of declining nominal house prices on listing, pricing, and time on the market in Boston operated through equity constraints. The estimates in Chan (2001) also suggested some nominal loss aversion in the New York area.

III. Data and Preliminary Analysis

The data used in the current paper are unique and are described in detail in the appendix. The primary data cover the 1985-96 period and are from the public-use version of the National Longitudinal Survey of Youth (NLSY79). These have been supplemented by data from three additional sources. First, administrative data on addresses were used to construct mobility histories for each household. This could not be done with data on the public-use and restricted-access Geocode files. Second, housing value and mortgage data from the NLSY79 master file were used. These data are top-coded in the public use version. Finally, Fannie Mae/Freddie Mac weighted-repeat-sales metropolitan house price indices were matched to each household-year observation. The final sample consists of 6,461 household-year observations. Sample descriptive statistics are given in Table A-1 in the appendix.

This sample has a number of advantages. First, it focuses on young households, many of whom own their first home. They are the most mobile and the most leveraged, and hence the most likely to be equity-constrained when house prices decline. This sets up a strong empirical test for equity constraints. If there is little evidence in favor of constraints in this sample, then equity is likely unimportant for mobility, because one would not expect it to affect the mobility of older, wealthier households. Alternatively, if there is evidence in favor of constraints, then equity is important for mobility, at least for young households.

Second, the sample spans a period of substantial variation in metropolitan housing market performance that can be used to identify any equity and loss aversion effects. This is particularly important for two reasons. First, the model and empirical work of Stein (1995) and Lamont and Stein (1999) on equity constraints has indicated the possibility of multiple equilibria, such that equity constraints may be at play in some metropolitan areas, but not others. Second, while both the empirical tests by Genesove and Mayer (1997, 2001) and Chan (2001) for equity constraints and loss aversion used incredibly rich data and have provided very persuasive evidence in favor of these phenomena, in the end, they apply only to the experiences of the condominium market in downtown Boston and the greater New York metropolitan area housing market. A question of first-order importance is whether these results generalize to the nation as a whole. This is especially important if, as outlined in the introduction, a main concern is that housing markets constrain mobility between metropolitan areas. Studies of single metropolitan areas cannot address this broader economic question.

Third, there are standard omitted variables issues. Previous studies have not been able to track changes in the demographic and economic circumstances of the households under study that almost surely affect mobility behavior strongly and may happen to be correlated with local housing and labor market conditions.[10] These include divorce, unemployment, and family decisions. Thus, another key question is whether once one controls for these other factors, evidence of equity constraints and loss aversion remains.

From the discussion in the previous section, there are many requirements for equity constraints to bind. First, most wealth must be in housing. If the household in the example above had $10,000 in other wealth and house prices fell 10 percent, the household would not be constrained because it could use its other wealth to make a down payment. Empirically, highly leveraged households are predominantly young, first-time owners with little other wealth. Table 1 shows the fraction of wealth in housing at the time of first home purchase for a sample of NLSY79 households. Overall, 80.5 and 90.6 percent of liquid assets go into housing at purchase, at the mean and median, respectively. Column (3) indicates that 25 percent of first-time homebuyers essentially have no other wealth after first purchase. These figures are higher when wealth is measured as liquid assets less debts.

Second, periods of declining nominal house prices are required for binding equity constraints, and, obviously, for nominal loss aversion.[11] Table 2 summarizes recent episodes of falling prices for selected United States metropolitan areas (grouped geographically). These episodes were measured using the Freddie Mac/Fannie Mae price indices described above. Price declines have been quite large. In Houston, there was a 27 percent decline from the market peak to the trough (Column 3). Peak-to-trough declines of 15 percent or more occurred in Texas, New England, and Southern California. These declines were large enough to have constrained most owners with less than 20 percent equity at the market peak.

Third, house prices must decline sufficiently rapidly for equity constraints to bind. If not, forward-looking owners could increase saving to maintain the option of purchasing another home.[12] The market peaks and troughs indicated in Columns (1) and (2) of Table 2 show that the duration of price declines has varied greatly. For example, the declines in Texas and New York lasted 2 to 3 years. In contrast, prices declined for five and seven years in California and Connecticut, respectively. To control for declines of different duration, Column (4) presents the annual average decline from peak to trough. By this metric, house prices declined rapidly, an annual average rate of 4.5-10 percent in the energy states, for example. Because many owners have 10 percent equity at purchase, these declines may have constrained a large number of households within one year.[13]

Table 3 examines the effect of falling nominal house prices on the distribution of housing equity in the NLSY79 sample. Panel A, Column (1) describes equity in the purchase year for all observations: 55.2 percent of owners had 20 percent or more in equity, 20.5 percent had between 10 and 19 percent equity, and 24.3 percent had less than 10 percent equity. Viewing the spell data as an (unbalanced) panel, Column (2) shows contemporaneous equity. As expected, equity grows over time: 63.1 percent of owners had 20 percent or more, 23.5 percent had between 10 and 19 percent, and just 13.4 percent had less than 10 percent.

Equity growth can come from a decline in mortgage debt through normal repayment, full or partial prepayment, and house price appreciation. Columns (3) and (4) repeat the tabulations in Columns (1) and (2) for the metropolitan areas with stable or rising housing markets, respectively. Columns (5) and (6) are defined similarly for the metropolitan areas with weak housing markets. A comparison of Columns (3) and (4) to (5) and (6) clearly illustrates the effect of appreciation. In falling markets, the distribution of equity shifts toward low equity between the initial and current periods. In stable and rising markets, the distribution shifts toward greater equity. Rising prices confer positive shocks; falling house prices confer negative shocks.

Because housing is the largest component of non-pension household wealth, house price fluctuations can have a large impact on the overall household balance sheet. Column (1) in panel *B* describes net worth relative to home value in the purchase year for all observations: 81.4 percent had net worth of 20 percent or more, 10.7 percent had between 10 and 19 percent, and 7.9 percent had less than 10 percent. Column (2) shows the contemporaneous ratio. As expected, net worth grows over time: 84.8 percent had 20 percent or more, 8.0 percent had between 10 and 19 percent, and just 7.2 percent had less than 10 percent. A comparison of Columns (3) and (4) with (5) and (6) clearly illustrates the effect of appreciation. In weak markets, the distribution of net worth remains very close between the initial and current periods. In stable and rising markets, the distribution shifts toward more net worth. Again, weak markets confer negative shocks to the whole balance sheet.

IV. Hazard Model Specification

To test for the presence of equity constraints and nominal loss aversion more formally, proportional hazard models of mobility are estimated. The 6,461 household-year observations described in the sample above comprise 3,112 residence spells. The sample has no left-censored spells. The longest (right-censored) spell is 12 years. Because intra- and intermetropolitan moves, as well as transitions between home owning and renting, have different determinants (Boehm, Herzog, and Schlottmann 1991; Bartel 1979), three possible transitions are studied: intrametropolitan moves from own-to-own, intrametropolitan moves from own-to-rent, and intermetropolitan moves. In particular, intrametropolitan moves typically are to adjust housing consumption; intermetropolitan moves typically are in response to employment changes.[14] Of the 3,112 spells, 895 were completed (i.e., ended in a move): 596 intrametropolitan own-to-own, 174 intrametropolitan own-to-rent; and, 125 intermetropolitan, respectively.

Figure 1 shows the Kaplan-Meier empirical hazards for each type of move. The hazard for intrametropolitan own-to-own moves rises and falls across spell periods, with peaks at almost 15 percent in year five and 4 percent in year eight. The shape of this hazard is similar to those in Sinai (1997), who used the Panel Study of Income Dynamics (PSID), and Chan (2001), who used Chemical Bank mortgage records. This is not inconsistent with young homeowners making lifecycle adjustments to housing consumption by trading up, perhaps to accommodate an increase in family size (Henderson and Ioannides 1989). The hazards for the other types of moves are remarkably flat across spell periods.

The econometric analysis employs a proportional hazard model of mobility, defined as a transition out of the current residence spell. Let i index households and t spell periods; then the hazard, λ, which measures the probability of moving in period t conditional on having not yet moved, is

$$\lambda_{it} = \lambda_{0t} \exp(Z_{it}{}' \beta) , \tag{1}$$

where λ_0 is the baseline hazard and Z is a vector of explanatory variables.

The household's home equity stake is measured by the loan-to-value (*LTV*) ratio. A high *LTV* means low equity. Mortgage underwriting guidelines suggest that households with an *LTV* greater than 0.80 (i.e., less than 20 percent equity) might be constrained. Mortgages with less than 20 percent down require the purchase of private mortgage insurance (PMI), which is expensive, 0.75 percentage points applied to the entire mortgage, not just the increment of the mortgage that would bring the down payment up to the 20 percent level.[15] Initially, all households with an *LTV* greater than 0.80, or less than 20 percent in equity, are considered constrained in the empirical analysis. This is the same definition used by Genesove and Mayer (1997, 2001), Lamont and Stein (1999), Caplin, Freeman, and Tracy (1997), and Chan (2001). Later, robustness checks will be performed with an *LTV* of 0.90 to define constrained households.[16]

Following Genesove and Mayer (2001), the basic specification is:

$$Z_{it}{}' \beta = X_{it}{}' \alpha + \theta(D_{it}^{LTV>0.80} \times LTV_{it}) + \gamma(D_{it}^{LOSS} \times LOSS_{it}) + \delta(D_{it}^{GAIN} \times GAIN_{it}) , \tag{2}$$

where $D_{it}^{LTV>0.80}$ is a dummy variable that equals one if *LTV* was greater than 0.80 and zero otherwise, and indicates whether the household is in the constrained group; D_{it}^{LOSS} is a dummy variable that is one if the household has experienced a nominal loss in house price in current period t relative to the initial purchase period 0; $LOSS$ measures this nominal loss in percentage terms and expresses it as a *positive* number (e.g., for a 5 percent nominal loss, $LOSS$ takes on a value of 0.05); D_{it}^{GAIN} is a dummy variable that is one if the household has experienced a nominal gain in house price in current period t relative to the initial purchase period 0; $GAIN$

measures this nominal gain in percentage terms (e.g., for a 5 percent nominal gain, $GAIN$ takes on a value of 0.05); and X is a vector of other explanatory variables.

The parameters in Equation (2) are interpreted as follows. First, θ is the effect of being equity constrained (conditional on X) on the hazard of moving. If constrained households have a lower hazard and, hence, a longer duration, then $\theta < 0$. Therefore, the simplest test for equity constraints implies a null hypothesis of $\theta = 0$. For intrametropolitan own-to-own moves, the alternative hypothesis is $\theta < 0$, i.e., lock-in. Importantly, because of the interaction of $D^{LTV>0.80}$ and LTV, Equation (2) says that it is not just whether the household has a high loan-to-value ratio (low equity), but also the *magnitude* of the equity that matters, so that $\theta < 0$ means that the larger the LTV (lower the equity) above the 0.80 threshold, the greater the reduction in the hazard of moving.[17] For intrametropolitan own-to-rent moves, the alternative hypothesis is $\theta > 0$, (i.e., switch tenure to mitigate the effect of the constraint). For intermetropolitan moves, the alternative is somewhat ambiguous, especially if most intermetropolitan moves are job-related and effectively exogenous to local housing market conditions. If low equity constrains these moves, then the alternative is $\theta < 0$. Alternative explanations may make $\theta > 0$ more plausible, e.g., if poor local economic conditions result in declining house prices, equity constraints, and, in response, constrained households more vigorously pursue opportunities in other metropolitan areas.

Second, γ is the effect of the size of the nominal loss on the hazard of moving for households with nominal losses. Therefore, under the null hypothesis of no nominal loss aversion, $\gamma = 0$. With the alternative hypothesis of loss aversion, $\gamma < 0$, which implies that the larger the loss, the lower the hazard of moving, and the longer the duration in the current residence.[18] Third, δ is the effect of the size of the nominal gain on the hazard of moving for households with nominal gains. Therefore, under the null hypothesis that nominal gains and losses are treated symmetrically, $\delta = -\gamma$ (because the loss variable takes on *positive* values).

Unfortunately, there are potentially serious econometric problems if observed *LTV* is used in Equation (2). First, observed *LTV* is likely endogenous. For example, a household that expects a long duration may be more likely to renovate, which would increase value and decrease *LTV*, ceteris paribus. In addition, if these renovations were financed with home equity borrowing, then mortgage debt would increase. More generally, any factor that affects the time path of mortgage debt may depend on expected duration, which would render *LTV* endogenous.

Second, there may be measurement error in mortgage debt and house value. In an interesting study, Goodman and Ittner (1992) used panel data on housing structures from the American Housing Survey and compared homeowners' self-reported estimates of the value of their homes with subsequent sale price of the home. They found that homeowners systematically overestimated the value of their home by 10 percent, but, somewhat surprisingly, this error was not correlated with any measured structural or household characteristics. Because the home value data in the NLSY79 are self-reported estimates, there is reason to believe there is reporting error in home values. In addition, a careful examination of the NLSY79 data shows that reported house and mortgage values were electronically miscoded from the interview information in some cases. The most typical coding error stems from the omission of the last digit of a reported home value or mortgage. For example, the same home worth $100,000 in one year is coded as being worth only $10,000 in the next year, followed by $100,000 in the subsequent year. Naturally, this means that in a large portion of the variation in *LTV* across time is due to measurement error.

To address endogeneity and measurement error, I define "simulated" *LTV*, denoted LTV^*. It is the loan-to-value ratio the household would have if it made regular payments on a 30-year fixed rate mortgage, never refinanced, and received the average metropolitan area appreciation (measured by the Freddie/Fannie indices). Specifically, simulated *LTV* for household i in metropolitan area *a* in spell period t is

$$LTV_{it}^{*} = \frac{M_{i0}}{V_{i0}} \times \frac{1}{I_{at}} \times \frac{[(1+r_0)^T - (1+r_0)^t]}{[(1+r_0)^T - 1]}, \tag{3}$$

where M_0, V_0, and r_0 are the mortgage principal, interest rate, and house value in the year of purchase. M_0 / V_0 is just the initial loan-to-value. I is an index of cumulative nominal house price appreciation in the metropolitan area between the purchase year (i.e., beginning of the spell), time 0, and the current spell period, time t. It has a value of 1 in the purchase year. If $I < 1$, then nominal house prices have fallen since purchase, which will cause simulated *LTV* to rise. The opposite holds for $I > 1$, which represents nominal housing appreciation. The variable r_0 is assumed to be the nominal 30-year fixed rate mortgage interest rate in the year of purchase. *T* is the term of the mortgage, assumed to be 30 years. The factor

$$\frac{[(1+r_0)^T - (1+r_0)^t]}{[(1+r_0)^T - 1]} \tag{4}$$

represents the rate of amortization of the mortgage principal, M.

Simulated *LTV* can be thought of as an instrumental variable. Because it is based on initial *LTV*, it is constructed to purge any endogeneity in the subsequent time paths of mortgage balance and house value that could affect the numerator and/or denominator of the observed *LTV*. It is highly correlated with observed *LTV* (the sample correlation coefficient is 0.86). In addition, it varies independently across individuals in the sample because households with different loan-to-value at purchase experienced different metropolitan area appreciation. Under the assumption that each household is a price-taker in the housing market, variation in metropolitan area appreciation is exogenous to a given household's mobility decision. Furthermore, simulated *LTV* is based on the loan-to-value in the year of purchase. A reporting error by the survey respondent with regard to total mortgage debt and house value in the purchase year is unlikely because the household went through the mortgage application and closing process prior to the NLSY79 interview for that year. In addition, by construction, simulated *LTV* will be uncorrelated with electronic coding error in the NLSY79 in the years subsequent to the purchase year. Hence, simulated *LTV* circumvents measurement error.

In the specifications reported below, simulated *LTV* is used in place of observed *LTV* in Equation (2). Thus, the specifications should be thought of as reduced-form relationships. Hereafter, the term "*LTV*" will refer to simulated *LTV* just described.

V. Estimation Results

The semiparametric estimator of Prentice and Gloeckler (1978) and Meyer (1988, 1990) is used, which models the baseline hazard flexibly as a vector of dummy variables, one for each spell period. Estimation results for intrametropolitan moves from own to own are shown in Table 4. Column 1 shows parameter estimates for Equation (2) with no household demographic or financial characteristics. Standard errors are in parentheses. Here, $\hat{\theta}$ is -0.166. The *p*-value for the test of the null hypothesis of no equity constraint versus the alternative of equity constraint is 0.049 and is shown in square brackets under the standard error. Because of the proportional hazard specification, this parameter estimate implies that binding equity constraints shift the baseline hazard downward by 13.9 percent when evaluated at the sample mean *LTV* for those in the constrained group. This is shown at the bottom of the table. Specifically, the sample mean *LTV* for those households in the constrained groups (i.e., with *LTV* greater than 0.80) was 0.905. The estimated percentage shift in the baseline hazard was calculated as

$$[\exp(\hat{\theta} \times \overline{LTV}^{C})] - 1, \qquad (5)$$

where $\overline{LTV}^{C}$ is the sample mean *LTV* for the sub-sample with $D^{LTV>0.80} = 1$, i.e., the constrained group.

Next, $\hat{\gamma}$ is -5.010. The *p*-value for the test of the null hypothesis that losses have no effect on mobility versus the alternative that losses reduce mobility as implied by nominal loss aversion is 0.0001. Therefore, the null hypothesis can be rejected in favor of loss aversion. This parameter estimate implies that losses shift the baseline hazard downward by 41.6 percent when evaluated at the sample mean loss for those with losses of 0.054 (i.e., a 5.4 percent nominal loss). Also, $\hat{\delta}$ is -9.872 and, based on the standard error, is significantly different from zero at the 0.01 percent level. This implies that households who experienced gains had lower hazards of intrametropolitan own-to-own moves and longer residence spells. When evaluated at the sample mean gain for those with gains of 0.135 (i.e., a 13.5 percent nominal gain), this estimate implies that gains shift the baseline hazard downward by 49.6 percent. Therefore, households hold on to gains. The fact that households hold on to both gains and losses suggests that households treat gains and losses asymmetrically.[19] This is confirmed by the formal statistical test of the null hypothesis that $\delta = -\gamma$, the *p*-value for which is 0.0001 and shown at the bottom of the table. Hence, the hypothesis can be rejected in favor of the alternative of asymmetric treatment of gains and losses, consistent with nominal loss aversion. Finally, a comparison of the estimated effects at the bottom of the table indicates that loss aversion is about three times (i.e., 41.6/13.9=2.99) more important than equity constraints in affecting intrametropolitan own-to-own mobility.

It is very likely that there are other characteristics of households that may be correlated with the loan-to-value and nominal loss that influence the hazard of moving. Therefore, in Column (2), a vector of demographic characteristics is added. These variables include dummy variables for whether the household is married, divorced, not employed, and white, respectively; dummy variables for whether the household head is 30 to 34 and 35 to 39 years old, respectively; and dummy variables for educational attainment.[20] Finally, there are dummy variables for whether there are children of various ages: age 5 or under, 6 to 10, 11 to 18, and more than 18 years old, respectively. These dummies are particularly important. Households with school-age children are thought to be less mobile.

Column (2) of Table 4 shows parameter estimates for Equation (2) with the demographic characteristics. Now, $\hat{\theta}$ is -0.150. The *p*-value for the test of the null hypothesis of no equity constraints versus the alternative of equity constraints is 0.068. In addition, there is still statistically significant evidence of loss aversion. Its effect on mobility is still about three times that of equity constraints. As expected, households with pre-school-age children have higher mobility. Given the importance of the quality of public schools in location decisions in a metropolitan area, these households may be moving to get into a suitable school district. Older, married, and more educated households have lower hazards of moving to another owner-occupied house in the same metropolitan area.[21]

The equity-constraints hypothesis focuses on the ability of the household to make a down payment on a desired home. But households contemplating a move must also be able to meet the flow cost of housing services out of their income. Although the hypothesis does not address this directly, it is important to control for it in the estimation. Therefore, in Column (3), a dummy variable that is one if the household would be "housing expenditure-constrained" and zero otherwise appears. This variable is meant to capture the flow cost of housing services relative to income. This is measured as the user cost of owner-occupied housing multiplied by house value, then divided by income. The household is considered constrained if this flow cost is more than one-third of household income.[22] The construction of this measure is described in detail in the appendix. Real income and net worth are added to the specification as well.

Column (3) shows parameter estimates for Equation (2) with the demographic and financial characteristics. Now, $\hat{\theta}$ is -0.132. However, the *p*-value for the test of the null hypothesis of no equity constraints is 0.105, so that the finding of equity constraints is no longer significant at the 5 percent level. In contrast, there continues to be statistically significant and economically important evidence of nominal loss aversion. Not surprisingly, households who currently are housing expenditure-constrained have significantly lower hazards of moving to another owner-occupied home in the same metropolitan area. The specification in Column (3) also controls for income and net worth directly. Conditional on the dummy for being expenditure constrained and the demographic characteristics, it does not appear that income and wealth have *independent* influences on the hazard of this type of move.[23]

One criticism of the specification in Equation (2) is that the implications for nominal loss aversion are drawn from the *GAIN* and *LOSS* variables that may just reflect local labor market conditions (Chan 2001). For example, metropolitan areas with good local labor markets also have stable or rising nominal house prices, so the finding that households hold on to gains just may reflect the fact that households with gains are in areas with good labor market opportunities that are valued enough so that households stay in those areas. Therefore, to insure that the estimates for the *GAIN* and *LOSS* parameters are not contaminated by unaccounted local labor market conditions, Column (4) includes the unemployment rate for the county of residence as an explanatory variable. This has the effect of reducing the parameter estimates for the *GAIN* and *LOSS* variables, but by a relatively small amount. Again, there is relatively weak statistical evidence for equity constraints but strong evidence for nominal loss aversion. Loss aversion continues to be about three times more important than equity constraints in its impact on mobility.

In Table 4, an *LTV* greater than 0.80, or less than 20 percent in equity, was used to define the constrained group. This follows Genesove and Mayer (1997, 2001), Lamont and Stein (1999), Caplin, Freeman, and Tracy (1997), and Chan (2001). However, with the increased promotion of mortgages with down payments of less than 20 percent in the sample period, it may be that many households with low equity were not actually constrained. Thus, Table 5 repeats the specifications in Table 4 for intrametropolitan, own-to-own moves, but defines low-equity households as those with a loan-to-value of 0.90 or higher. This is a stricter measure of the equity constraint than is used by others in the literature. Qualitatively, the results are quite similar. There is statistically significant evidence of equity constraints, but its precision weakens as more explanatory variables are added.[24] In addition, there is significant evidence of nominal loss aversion, and nominal aversion dominates equity constraints in impact by a ratio of about 2.5:1.

Table 6 gives estimation results for the same specifications in Table 4 but for intrametropolitan, own-to-rent moves. Based on the *p*-values shown in the first row of the table, the null hypothesis of no equity constraints for moves from own to rent in a metropolitan area cannot be rejected. This is true for all specifications. In addition, once the additional demographic and financial covariates are accounted for, there is no statistically significant effect of nominal losses on the hazard of this type of move (Columns (2) through (4)). Consistently, being white, married, more educated, older, wealthier, not housing-constrained, and having experienced nominal housing gains all reduce the likelihood of an own-to-rent move. The age of children does not matter for intrametropolitan own-to-rent mobility.

Table 7 gives estimation results for the same specifications in Tables 4 and 6 but for intermetropolitan moves. Based on the *p*-values shown in the first row of the table, the null hypothesis of no equity constraints for moves from own to rent in a metropolitan area cannot be rejected. This is true for all specifications. Nor is there evidence for loss aversion. Consistently, being married and older reduce the likelihood of an intermetropolitan move, while having a college degree and being unemployed increase the likelihood of such a move. The facts that equity constraints do not hinder intermetropolitan mobility and that those not employed are more likely to move between metropolitan areas cast some doubt on the hypothesis that housing markets retard labor migration in the United States (Oswald 1996, 1997, 1999).

VI. Conclusion

The empirical analysis yielded five principal findings. First, household intrametropolitan own-to-own mobility responds differently to nominal housing losses than to gains. These effects were estimated controlling for equity constraints, so that these results may be interpreted as evidence of nominal loss aversion. These findings confirm those in Genesove and Mayer (2001), and complement the results of Engelhardt (1996a), who found evidence of the asymmetric treatment of housing gains and losses on homeowner saving behavior. Second, nominal loss aversion is significantly less pronounced in intrametropolitan own-to-rent and intermetropolitan mobility, respectively. Third, there is some evidence of binding equity constraints in intrametropolitan own-to-own mobility (although it is weaker in a statistical sense than that for loss aversion). This complements the findings of Chan (2001) on mobility and Caplin, Freeman, and Tracy (1997) on mortgage refinancing. Fourth, there is little evidence that low equity constrains intrametropolitan own-to-rent and intermetropolitan mobility, respectively. Fifth, a comparison of the estimated effects indicates that nominal loss aversion has a more dominant effect than equity constraints in restricting household mobility:,roughly two and- one-half to three times the impact of equity constraints. This result complements the finding of Genesove and Mayer (2001) for the downtown Boston condominium market.

The labor-market implications are somewhat mixed. First, the fact that intrametropolitan own-to-rent and intermetropolitan moves are not constrained suggests that declining housing equity due to falling nominal prices does not impede regional labor market adjustment necessarily. This is inconsistent with Oswald (1996, 1997, 1999), but not Blanchard and Katz (1992). On the other hand, because intrametropolitan moves for those who wish to remain homeowners are constrained, there might be some locationally mismatched workers within a metropolitan area. The magnitude of this potential welfare loss is unclear, and its estimate is far beyond the scope of this paper. Nonetheless, if intrametropolitan moves are primarily to adjust housing consumption, then constrained households are worse off from declining prices, but there will be little labor market impact. Alternatively, if such moves are contemplated to take employment in another part of the metropolitan area, to which commuting from the current residence would be prohibitively costly, then there could be some labor market distortions. Some of the literature on joint housing, employment, and commuting choice suggests that homeowners are more willing to take on a longer commute for a new job than to change residences within a metropolitan area. This suggests that intrametropolitan labor market distortions from declining equity may not be large. This is a clear avenue for future research.

Appendix

This appendix describes the construction of the data set. The primary data cover the 1985-96 period and are from the 1985-1998 waves of the National Longitudinal Survey of Youth (NLSY79). Engelhardt (1998) and Zagorsky (1999) discussed the quality of these data. Haurin, Hendershott, and Kim (1994) and Haurin, Hendershott, and Wachter (1996, 1997) have used these data to analyze housing decisions of young households. The NLSY79 started as a national, stratified, random sample of 14 to 21 year-olds in 1979. The survey was conducted every year from 1979 to 1994; after 1994, it was conducted every two years. It asked detailed questions about education, employment, income, home ownership, family background, etc. In 1985, questions about assets and debts were added, including mortgage debt and home value. Because home ownership status has been asked each year since 1985, it is possible to completely track residence transitions from early adulthood. As a result, the sample has no left-censored spells. This means that the earliest a spell could have begun and been included in the sample is 1985. In turn, the longest spell observable with these data is 12 years. A potential criticism of this sample is that the distribution of completed residence spells may be poorly estimated if the average spell length of homeowners of this age is greater than 12 years, the maximum spell length in the sample. Sinai (1997) studied housing mobility for homeowners of all ages in the 1970-91 waves of the Panel Study of Income Dynamics (PSID). His analysis showed that the hazards for homeowners declined steadily for spells of 10 years or less. For spells greater than 10 years, the hazards were roughly flat for all types of transitions. In addition, he estimated the average duration for a homeowner as 6.8 years. Under the assumption that young homeowners (age 20 to 41 in my sample) have shorter completed spells than the average-aged homeowner in Sinai's analysis—which seems plausible—then the NLSY79 sample may estimate the underlying spell distribution well. In addition, Sinai presented sensitivity analyses that showed that truncation of the spell length at 8 years had little effect on the hazard estimates.

Mobility Histories

The empirical analysis focuses on three possible transitions: intrametropolitan moves from own to own, intrametropolitan moves from own to rent, and intermetropolitan moves. Unlike other panel studies, such as the *Panel Study of Income Dynamics* (*PSID*), the NLSY79 did not ask a question each year about whether the respondent had moved since the previous interview. The public-use version of the NLSY79 has information on home ownership in each year. Hence all own-to-rent moves can be tracked in this data set. The restricted-access Geocode dataset gives information on state, county, and the metropolitan area of residence. When combined, the public-use and Geocode data can track own-to-rent and intercounty moves. Unfortunately, intracounty own-to-own moves cannot be tracked. Because most moves are local, and most metropolitan areas are comprised of just a few (and, in some cases, one) counties, the combined public-use and Geocode data significantly understate the number of actual transitions. To overcome this problem, I obtained permission from the United States Department of Labor, Bureau of Labor Statistics, to use administrative address data on the NLSY79 respondents in each survey year to construct a mobility history for each respondent. These mobility data were provided graciously by Patricia Reagan, who assembled them at the Center for Human Resources Research at the Ohio State University. When comparing the county of residence from the address records to that in the Geocode file, a number of errors were found in the Geocode data. All state and county codes used in this study were based on the administrative address records.

Metropolitan House Price Indices

Because the empirical analysis focuses on the effect of house prices on mobility, only respondents in metropolitan areas with available house price information were included in the sample. Metropolitan house prices were measured by the Fannie Mae/Freddie Mac Weighted Repeat Sales Price index. This index is discussed in detail in Abraham and Hendershott (1992). This is available in the 1985-96 period for the following 149 metropolitan areas: Akron, OH; Albany, NY; Albuquerque, NM; Allentown, PA; Ann Arbor, MI; Appleton, WI; Atlanta, GA; Atlantic-Cape May, NJ; Augusta, GA; Austin, TX; Bakersfield, CA; Baltimore, MD; Barnstable-Yarmouth, MA; Baton Rouge, LA; Bellingham, WA; Birmingham, AL; Bloomington-Normal, IL; Boston, MA; Boulder-Longmont, CO; Bridgeport, CT; Brockton, MA; Buffalo, NY; Burlington, VT; Canton-Massillon, OH; Cedar Rapids, IA; Charleston-North Charleston, SC; Charlotte, NC; Chicago, IL; Cincinnati, OH; Cleveland, OH; Colorado Springs, CO; Columbia, SC; Columbus, OH; Dallas, TX; Danbury, CT; Davenport-Moline-Rock Island, IL; Dayton-Springfield, OH; Daytona Beach, FL; Denver, CO; Des Moines, IA; Detroit, MI; Eugene-Springfield, OR; Evansville-Henderson, IN-KY; Flint, MI; Fort Collins-Loveland, CO; Fort Lauderdale, FL; Fort Wayne, IN; Fort Worth-Arlington, TX; Fresno, CA; Gary, IN; Grand Rapids-Muskegon-Holland, MI; Green Bay, WI; Greensboro-Winston-Salem-High Point, NC; Greenville-Spartanburg, SC; Hamilton-Middletown, OH; Harrisburg-Lebanon-Carlisle, PA; Hartford, CT; Houston, TX; Huntsville, AL; Indianapolis, IN; Jacksonville, FL; Kalamazoo-Battle Creek, MI; Kansas City, MO; Knoxville, TN; Lancaster, PA; Lansing-East Lansing, MI; Las Vegas, NV; Lawrence, MA-NH; Lexington, KY; Lincoln, NE; Little Rock-North Little Rock, AR; Los Angeles, CA; Orange County, CA; Riverside, CA; Louisville, KY; Lowell, MA-NH; Madison, WI; Melbourne-Titusville-Palm Bay, FL; Memphis, TN-AR-MS; Miami, FL;

Milwaukee-Waukesha, WI; Minneapolis-St. Paul, MN; Modesto, CA; Monmouth-Ocean, NJ; Nashua, NH; Nashville, TN; New Haven-Meriden, CT; New Orleans, LA; New York, NY; Bergen-Passaic, NJ; Newark, NJ; Middlesex-Somerset-Hunterdon, NJ; Nassau-Suffolk, NY; Norfolk-Virginia Beach-Newport News, VA; Oklahoma City, OK; Omaha, NE; Orlando, FL; Peoria, IL; Philadelphia, PA; Phoenix-Mesa, AZ; Pittsburgh, PA; Portland, ME; Portland-Vancouver, OR-WA; Portsmouth-Rochester, NH-ME; Providence-Fall River, RI-MA; Provo-Orem, UT; Racine, WI; Raleigh-Durham, NC; Reading, PA; Reno, NV; Richmond, VA; Rochester, NY; Rockford, IL; Sacramento, CA; Saginaw-Bay City-Midland, MI; St. Louis, MO; Salem, OR; Salinas, CA; Salt Lake City-Ogden, UT; San Antonio, TX; San Diego, CA; San Francisco, CA; Oakland, CA; San Jose, CA; San Luis Obispo, CA; Santa Barbara, CA; Santa Cruz, CA; Santa Rosa, CA; Sarasota-Bradenton, FL; Seattle-Bellevue-Everett, WA; Springfield, IL; Springfield, MA; Stamford-Norwalk, CT; Stockton, CA; Syracuse, NY; Tacoma, WA; Tampa-St. Petersburg-Clearwater, FL; Toledo, OH; Trenton, NJ; Tucson, AZ; Tulsa, OK; Vallejo, CA; Ventura, CA; Visalia-Porterville-Tulare, CA; Washington, DC-MD-VA; West Palm Beach-Boca Raton, FL; Wichita, KS; Wilmington-Newark, DE-MD; Worcester, MA; and York, PA.

Income

The household income measure used is real total net family income (in 1993 dollars). Interviews typically were conducted in the spring of the calendar year. The survey asked about income earned in the previous calendar year. For example, the 1993 wave contains information on 1992 income. This means that the 1986 to 1994 surveys provide information on income in calendar years 1985 to 1993. After 1994, the survey went to an every-other-year format, but questions on income still referred to the previous calendar year. This means that the 1996 survey year gives information on calendar year 1995 income, and the 1998 survey year gives information on calendar year 1997 income. Incomes from calendar years 1994 and 1996 were not asked. Therefore, for this study, income for calendar years 1995 and 1997 proxy for those in 1994 and 1996, respectively. All income values in the paper are in real 1993 dollars, deflated by the All-Items CPI.

House Value and Mortgage Data

The public use NLSY79 top-coded housing value and mortgage debt at $150,000 in nominal terms for the 1985 to 1994 waves (Engelhardt 1998). Initial loan-to-value cannot be calculated for observations with top-coded values, and these observations must be excluded from the sample. In 1985, less than 2 percent of observations had top-coded values for house value and mortgage debt. But because the top-code threshold was fixed in *nominal* terms, over time with inflation, a growing fraction of observations had top-coded values: 17.38 percent for housing value and 6.35 percent for mortgage debt in 1994, respectively. However, because of the substantial regional variation in house price levels, the truncated cases came disproportionately from high-cost markets such as Boston, San Francisco, Los Angeles, and New York. Furthermore, these markets experienced steep declines in nominal house prices in the late 1980s and early 1990s. Therefore, the exclusion of observations with top-coded values results in differential sample selection and potentially biased estimated equity effects. To overcome this problem, I obtained permission from the United States Department of Labor, Bureau of Labor Statistics, to use housing value and mortgage data from the NLSY79 master file at the Center for Human Resources Research (CHRR) at the Ohio State University. Specifically, CHRR released to me new house value and mortgage data in which the top-code thresholds were adjusted upward so that only approximately 2 percent of the observations in each of the 1985 to 1998 survey years had top-coded house values and mortgage amounts. These new data were used in the empirical analysis. All asset and debt values in the paper are in real 1993 dollars, deflated by the All-Items CPI.

Assets and Debts

For budgetary reasons, questions on assets and debts were not asked in the 1991 wave of the NLSY79 (Engelhardt 1998). However, questions about income from assets in 1991 were asked in the 1992 wave, and asset and debt questions were asked in the 1990 and 1992 waves. So, for 1991 the asset income was capitalized at the prevailing annual return. This, along with information from 1990 and 1992, was used to impute assets and debts for each household in 1991. The empirical results were not sensitive to the exclusion of all 1991 observations.

Dummy If Housing Expenditure-Constrained

Let i index households and t index calendar years; then, following Poterba (1991), whether or not the household claims itemized deductions for mortgage interest and property taxes paid depends on the tax saving from itemizing, ξ,

$$\xi_{it} = \theta_{it}[\tau_{it}^{s} + (\tau^{p} + i_0 LTV_{it})V_{it}^{*}] - S_{it}, \tag{A1}$$

where τ^{p} is the property tax rate, i_0 is the nominal mortgage interest rate in the year of purchase, LTV is the loan-to-value rate, and S is the standard deduction amount. V^{*} is an exogenous measure of house value, similar in spirit to simulated loan-to-value. It is the household's house value if the home had appreciated at the average metropolitan rate, I_{at}, defined in the text. θ is the household's federal marginal tax rate on the first dollar of itemized deduction. The variable τ^{s} is the household's state marginal tax rate on the first dollar of itemized deduction. If $\xi_{it} \geq 0$, then the household will itemize and the marginal user cost of owner-occupied housing (as a fraction of the house price) is

$$u_{it}^{m} = (1-\theta_{it})(i_t + \tau_{it}^{s} + \tau^{p}) + d + a + m - \pi_t^{e}, \tag{A2}$$

where d is the physical rate of decay, m is maintenance expenditure, a is a risk factor, and π^{e} is expected appreciation. If $\xi_{it} < 0$, then the household will not itemize and the marginal user cost is

$$u_{it}^{m} = ((1-LTV_{it})(1-\theta_{it}) + LTV_{it})i_t + \tau^{p} + d + a + m - \pi_t^{e}. \tag{A3}$$

Following Poterba (1984, 1991), the user cost is calibrated for each household under the following assumptions: $\tau^{p} = 0.02$, $, d = 0.014$, $a = 0.05$; and i is the rate on a 30-year fixed rate mortgage. The federal and state tax first-dollar marginal tax rates were calculated using the NBER TAXSIM calculator. The dummy if housing expenditure-constrained was constructed to take on a value of 1 if the flow cost of housing relative to income,

$$\frac{u_{it}^{m} V_{it}^{*}}{y_{it}}, \tag{A4}$$

(where y is household income) was greater than 0.33 and zero otherwise. A key assumption in calculating (A4) is what *LTV* to use in (A1) and (A3). Variants of this variable were constructed using an *LTV* of 0.80, 0.90, 0.95, as well as contemporaneous simulated *LTV*; the estimation results were remarkably robust across these alternative specifications. The results in Tables

4 through 7 in the paper used an *LTV* of 0.95, which is akin to assuming that the household would take out a mortgage on the next home with just 5 percent down. Such mortgages were prevalent by the end of the sample period. This assumption helps insure that the variable really picks up expenditure-constrained households, for these would be households who could not buy back their current residence and spend less than one-third of their income on housing. Variants of this variable were also constructed using flow-cost-of-housing-to-income ratios of 0.25, 0.28, 0.30, and 0.40 to define the dummy. Again, the estimation results were quite robust to these alternatives.

Table A-1
Sample Means (Standard Deviations) for the Explanatory Variables

	(1)	(2)	(3)	(4)
		SubSample of Observations with		
Explanatory Variable	**Full Sample**	**Intrametropolitan Own-to-Own Moves**	**Intrametropolitan Own-to-Rent Moves**	**Intermetropolitan Moves**
Dummy if LTV>0.80	0.317	0.303	0.335	0.367
×LTV	(0.433)	(0.428)	(0.442)	(0.455)
Dummy if Nominal Loss	0.0052	0.0050	0.204	0.225
×Nominal Loss	(0.022)	(0.019)	(0.394)	(0.412)
Dummy if Nominal Gain	0.064	0.069	0.051	0.068
×Nominal Gain	(0.112)	(0.100)	(0.092)	(0.130)
Dummy if Married	0.799	0.803	0.730	0.784
Dummy if White	0.818	0.837	0.666	0.856
Dummy if Age 30 to 34	0.485	0.518	0.466	0.496
Dummy if Age 35 to 39	0.144	0.117	0.080	0.096
Dummy if Some College	0.258	0.272	0.270	0.168
Dummy if College Degree	0.318	0.322	0.121	0.520
Dummy if Children Age 5 and Under	0.432	0.477	0.454	0.400
Dummy if Children Age 6 to 10	0.330	0.312	0.419	0.304
Dummy if Children Age 11 to 18	0.185	0.178	0.230	0.144

Table A-1 (Continued)

Explanatory Variable	(1) Full Sample	(2) Sub-Sample of Observations with Intrametropolitan Own-to-Own Moves	(3) Intrametropolitan Own-to-Rent Moves	(4) Intermetropolitan Moves
Dummy if Children over Age 18	0.021	0.015	0.034	0
Dummy if Not Employed	0.172	0.179	0.183	0.240
Dummy if Divorced	0.067	0.083	0.109	0.056
Dummy if Housing Expenditure-Constrained	0.195	0.183	0.264	0.208
Real Income	62,173 (98,336)	66,017 (104,773)	57,746 (112,193)	91,347 (180,154)
Real Net Worth	67,736 (74,364)	70,503 (81,032)	42,387 (48,715)	75,666 (76,293)
County Unemployment Rate	6.23 (2.46)	6.08 (2.33)	6.35 (2.29)	6.37 (2.65)
Dummy if LTV>0.90 $\times$LTV	0.161 (0.360)	0.146 (0.347)	0.204 (0.394)	0.225 (0.412)
Number of Observations	6,461	596	174	125

Note: Sample means of the explanatory variables, with standard deviations for all continuous variables shown in parentheses. Income and net worth are in thousands of 1993 dollars. The county unemployment rate is measured in percentage points. The sample mean *LTV* for those in the constrained group was 0.905. The sample mean nominal loss for those who experienced nominal losses was 0.054, or a 5.4 percent loss. The sample mean nominal gain for those who experienced nominal gains was 0.135, or a 13.5 percent gain.
Source: Author’s calculations.

Endnotes

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1. Green and Hendershott (2001) have analyzed this in the United States as well. Some aspects of the link between housing and mobility have been explored by Bover, Muellbauer, and Murphy (1989), Hughes and McCormick (1981, 1985, 1987), Henley (1998a,b), Boheim and Taylor (2000), and Gardner, Pierre, and Oswald (2000) in the United Kingdon and Van Der Berg (1992) in the Netherlands.

2. The down payment requirement for conventional mortgages ranges from 10 to 20 percent. Changes in secondary mortgage market underwriting guidelines have made 5 percent down mortgages more prevalent.

3. Another possible source of down payment funds is transfers from family or friends. Engelhardt and Mayer (1998) found that about 25 percent of first-time buyers receive such transfers, but among repeat buyers they were rare (4 percent).

4. This example abstracts from other costs that may further deter households. These include private mortgage insurance (discussed later in the text), closing costs, broker costs, and moving costs. Moving costs can be at least \$750 to \$4500, depending on the area, distance, and volume. For a \$155,000 mortgage in the greater New York metropolitan area, Caplin, Freeman, and Tracy (1997) estimated the closing costs to be between \$5,100 and \$8,400, or 3.3 and 5.4 percent of the loan, respectively. In general, brokers are paid 6 percent of the sale price of the home.

5. It is important to note that this leverage effect is independent of a wealth effect from the capital gain (or loss). To illustrate this, consider a model with no mortgages—all homes must be purchased in cash. A household owning a $100,000 home who experienced 10 percent appreciation could afford to purchase just a $110,000 home. The fact that households can borrow to finance the purchase of a home means that a dollar of capital gain can buy more than a dollar of housing.

6. Stein (1995) referred to this as "packing." Mayer (1993) has provided striking evidence on this. He examined the level of housing leverage in Massachusetts (predominantly the greater Boston metropolitan area) during the boom and bust of the 1980s and 1990s. He found that of the 580,000 households who purchased single family homes between 1982 and 1992, more than 150,000 had less than 5 percent equity in 1992, and the majority of these had no equity. Even lower equity levels were found in the condominium market. These findings suggest the great potential for collateral constraints. Lamont and Stein (1999) observed substantial variation in packing across metropolitan areas using data from the American Housing Surveys.

7. In this light, housing equity plays a role similar to collateral for firms. There, temporary economic shocks that depress the value of assets used for productive purposes and collateral can reduce the net worth of firms, reduce the asset demand for constrained firms, and result in lower asset prices. This further reduces net worth and feeds back into prices. This link between asset prices and collateral has been examined recently by Kiyotaki and Moore (1997) and Shleifer and Vishny (1992), among others.

8. Kiyotaki and Moore (1997) provide a summary of the empirical evidence consistent with collateral constraints for firms. Henley (1998a) examined the effect of negative housing equity on the mobility of British households in the early 1990's. He used a sample of 3,530 households from the 1991 to 1994 waves of the British Household Panel Survey and estimated a semiparametric duration model of mobility with competing risks. The three risks modeled were a move to an owner-occupied home, a move to a public-sector rental, and a move to a private-sector rental. He found significant lock-in from fallen prices. His estimates suggested that the mobility of negative equity households would have been 50 percent higher if they had had a positive equity position.

9. Three related studies have found evidence that low equity constrains mortgage refinancing in the United States. Caplin, Freeman, and Tracy (1997) used a national sample of 35,865 Chemical Bank mortgages. They examined prepayments and found that constrained households were significantly less likely to refinance. Constrained households were defined as those with less then 20 percent equity in states with weak housing markets (Connecticut, Florida, Massachusetts, New Jersey, and Rhode Island). Peristiani et al.(1996) examined mortgage refinancing with a sample of mortgages from the Mortgage Research Group for a number of counties in five states. The counties examined were Los Angeles, Riverside, and Ventura Counties, CA; Citrus, Clay, Escambia, Hernando, Manatee, and Marion Counties, FL; Cook County, IL; Bergen, Essex, and Monmouth, NJ; and, Orange County, NY. They found that low equity and poor credit history significantly decreased refinancing. Finally, Archer, Ling, and McGill (1996) used a panel of 5,042 households from the 1985 and 1987 American Housing

Surveys. They estimated that constrained households (i.e., less than 10 percent equity) were significantly less likely to refinance.

10. For example, the LINK data used by Genesove and Mayer (1997) had very detailed information on the listing behavior of sellers in the downtown Boston condominium market, which were necessary for the very clean empirical tests for equity constraints on seller behavior. But these data do not provide information on motivations for sale (Glower, Haurin, and Hendershott (1998)). Chan (2001) used Chemical Bank mortgage records for the New York metropolitan area, but these data only provided detailed financial and demographic data at the time of underwriting. In both studies, nothing is known about what precipitated the move or the type of move (intra- vs. intermetropolitan or own-to-own vs. own-to-rent).

11. It is important to emphasize that equity effects require declining nominal prices. Real prices can fall even when nominal prices are flat or rising, as long as inflation exceeds nominal appreciation. In addition, loss aversion is with respect to nominal, not real, losses. Genesove and Mayer (2001) provide evidence that it is nominal losses that matter. Seller behavior is only weakly affected by real losses. In the analysis below, there was no evidence that real losses mattered.

12. This issue is not addressed in the static model of Stein (1995) but is in the dynamic model of Ortalo-Magne and Rady (1998). Engelhardt (1996a) examined the effect of housing gains and losses on homeowner saving behavior.

13. Furthermore, surveys of homeowners in boom and bust housing markets by Case and Shiller (1988) have provided provocative evidence that homeowners are not forward-looking. Rather, they seem to base their expectations of future price movements on past price behavior, not fundamentals, so that it may take a number of years of declining prices until households expect prices to decline and adjust their behavior. Based on the declines in Column 4, even two years of prices falling at an annual rate of 2.5 percent (roughly the lower bound in Column 4) would have been enough to have constrained highly leveraged owners.

14. In the sample, many intermetropolitan own-to-rent movers bought a home a year after arrival in the new metropolitan area. This is consistent with renting as a short-term, owner-occupied housing search device. Because the subsample of the remaining (i.e., “permanent”) intermetropolitan own-to-rent movers was so small, reliable estimates for this transition could not be made. Consequently, all intermetropolitan movers were pooled into one category.

15. In fact, for an individual purchasing a home with 15 percent down, the shadow rate of return on the additional down payment of 5 percent would be equal to the mortgage interest rate plus an additional 12 percentage points for PMI.

16. It would be very interesting to examine stricter measures of constraints, such as an *LTV* of 0.95, or even negative equity (Henley 1998a). However, the cell sizes become too small for reliable estimation and inference for these categories in this sample.

17. In principle, the effect of the nominal loss on low-equity households may be nonlinear if the loss is so large that the mortgage default option is in the money. This is not addressed in this paper because the cell sizes became small as loan-to-value rose above one, which resulted in unreliable estimation and inference in this sample. Moreover, there is an extensive literature on the effect of house price fluctuations on mortgage default, most of which uses data that are superior for default models than these data. Mattey and Wallace (1998, 1999) review much of this literature and discuss the role of house price declines on mortgage defaults.

18. Note that this is consistent with Genesove and Mayer (2001). If households are loss averse, they will attempt to attenuate losses by setting a higher list price and accept longer time on the market. But longer time on the market (including withdrawal of the property) implies a longer duration and a lower hazard of moving.

19. Remember that, in this context, symmetric treatment of gains and losses would be that households hold gains and shed losses, ceteris paribus.

20. The NLSY79 is a longitudinal survey that began in 1979 with 12,686 individuals between the ages of 14 and 21. By the 1996 interview data, the last year in my sample, these individuals were between the ages of 31 and 39. Hence, there are no individuals older than 39 in the sample, and that is why the age variables only go to 39 years. In addition, because the NLSY79 focused on a birth cohort, after conditioning on other covariates, there is little true age variation in the sample.

21. One interpretation is that these factors may proxy for financial security; and financially secure households may be better able to obtain a good housing match and, therefore, move less often.

22. Different cutoffs defining constrained households were used and did not change the empirical findings. Overall, the results were quite robust to changes in the definition of this variable.

23. Alternate measures of wealth were used to check the robustness of the estimates for this specification. These measures included non-housing wealth, financial wealth, financial assets, and highly liquid assets. These alternative specifications produced economically very similar and statistically significant lock-in effects to those shown in Column 3 of Table 4. In addition, specifications were estimated that used the measure of "extended *LTV*" from Chan (2001), which is defined as loan balance less other assets, divided by house value. The results were similar in economic magnitude and statistical significance to those shown here. Furthermore, a number of alternative measures of the housing expenditure constraint were specified. These are described in detail in the appendix. Again, all of these specifications produced estimates quantitatively similar to those presented here. These specifications are available upon request. Overall, these findings were very robust to alternative specifications. Finally, quadratic terms in the loss and gain variables were added to the model, but the null hypothesis of linearity in each could not be rejected at conventional significance levels.

24. The *p*-values for the test of the null hypothesis of no equity constraints in Table 5 are larger than those in Table 4 because the definition of constrained at an *LTV* of 0.90 generates smaller cell sizes for the constrained group, which reduces the precision in estimation and raises the standard errors.

Table 1. Percent of Wealth in Housing at First Home Purchase

Wealth Measure	(1) Mean	(2) Median	(3) 25th percentile	(4) 75th percentile
Liquid Assets	80.5	90.6	68.9	99.0
Liquid Assets Less Debts	83.6	94.4	72.7	100.0

Note: Total liquid assets is the amount of the down payment plus the value of financial assets in the form of savings accounts, money market deposit accounts, certificates of deposit, interest-earning checking accounts, United States saving bonds, individual retirement accounts (IRAs), 401(k)-type pension arrangements, Keogh plans, mortgages held by the household, non-interest earning checking accounts, money market funds, United States government securities, municipal and corporate bonds, stocks and mutual fund shares, money owed by others to the household, and other interest-earning-assets. Debts are the sum of farm, business, and non-owner-occupied real estate debt and non-vehicle-related debt.
Source: Author's calculations from the sample of all first-time homebuyers in the 1985-90 waves of the NLSY79

Table 2. Episodes of Falling Nominal House Prices for Selected Metropolitan Areas

Metropolitan Area	(1) Market Peak (Yr:Qtr)	(2) Market Trough (Yr:Qtr)	(3) Peak-to-Trough Total Decline (percent)	(4) Peak-to-Trough Annual Average Decline (percent)
Houston, TX	83:2	87:4	27.2	5.5
Dallas, TX	86:2	89:1	14.4	5.0
Austin, TX	86:2	88:4	26.9	10.0
San Antonio, TX	86:1	90:2	20.0	4.4
Oklahoma City, OK	86:2	88:3	22.7	9.5
Tulsa, OK	83:3	89:1	15.3	2.6
New Orleans, LA	86:2	88:4	11.9	4.6
Baton Rouge, LA	86:1	89:1	14.1	4.5
Denver, CO	86:2	89:1	6.8	2.4
Boston, MA	88:4	92:2	9.7	2.7
Portsmouth, NH	89:1	92:3	15.3	4.2
Providence, RI	89:4	94:4	11.8	2.3
Hartford, CT	88:3	95:1	19.7	2.8
New Haven, CT	88:2	95:1	20.8	2.8
New York, NY	89:1	91:3	6.4	2.5
Nassau-Suffolk, NY	88:3	91:2	9.0	3.2
Middlesex-Somerset, NJ	88:2	91:3	10.5	3.1
Bergen-Passaic, NJ	88:2	91:3	10.0	3.0
Los Angeles, CA	90:1	95:1	21.5	4.0
Orange County, CA	90:1	95:1	17.9	3.4
Riverside, CA	91:1	95:1	19.1	4.5
Santa Barbara, CA	90:3	95:1	12.4	2.6
San Diego, CA	90:3	95:1	10.2	2.2
San Francisco, CA	90:1	94:4	11.0	2.2
San Jose, CA	89:4	94:3	12.2	2.4

Source: Author's calculations using Freddie Mac/Fannie Mae weighted-repeat-sales quarterly house price indices for each of the metropolitan areas shown.

Table 3. The Effect of Falling Nominal House Prices on the Distribution of Housing Equity and Net Worth

	(1)	(2)	(3)	(4)	(5)	(6)
	Percent in Cell					
	All Observations		Observations in Stable or Rising Markets		Observations in Falling Markets	
Category	Purchase Year	Current Year	Purchase Year	Current Year	Purchase Year	Current Year
A. Housing Equity as a Percent Age of Home Value						
20 Percent or More	55.2	63.1	54.8	64.3	59.0	52.1
10-19 Percent	20.5	23.5	21.2	17.7	14.0	15.5
Less Than 10 Percent	24.3	13.4	24.0	18.0	27.0	32.4
B. Net Worth as a Percent Age of Home Value						
20 Percent or More	81.4	84.8	81.8	85.6	77.1	77.8
10-19 Percent	10.7	8.0	10.7	7.7	10.9	10.9
Less Than 10 Percent	7.9	7.2	7.5	6.7	12.0	11.3

Source: Author's calculations from the sample of 6,461 household-year observations described in the text.

Table 4. Semiparametric Proportional Hazard Model Estimates of Intrametropolitan Own-to-Own Mobility, with Loan-to-Value Threshold at 80 Percent

Explanatory Variable	(1)	(2)	(3)	(4)
Dummy if LTV>0.80 ×LTV	-0.166	-0.150	-0.132	-0.123
	(0.100)	(0.101)	(0.106)	(0.106)
	(0.049)	(0.068)	(0.105)	(0.123)
Dummy if Nominal Loss ×Nominal Loss	-5.010	-7.736	-7.544	-5.855
	(2.452)	(2.457)	(2.457)	(2.472)
	(0.0001)	(0.001)	(0.001)	(0.009)
Dummy if Nominal Gain ×Nominal Gain	-9.872	-3.693	-3.675	-3.455
	(0.529)	(0.570)	(0.572)	(0.575)
	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Baseline Dummy: Year 1	-3.730	-3.200	-3.180	-3.047
	(0.212)	(0.263)	(0.266)	(0.278)
Baseline Dummy: Year 2	-3.321	-2.841	-2.833	-2.727
	(0.215)	(0.260)	(0.263)	(0.274)
Baseline Dummy: Year 3	-2.922	-2.494	-2.483	-2.367
	(0.212)	(0.255)	(0.258)	(0.269)
Baseline Dummy: Year 4	-2.968	-2.516	-2.510	-2.409
	(0.235)	(0.275)	(0.277)	(0.287)
Baseline Dummy: Year 5	-2.656	-2.259	-2.250	-2.173
	(0.217)	(0.255)	(0.256)	(0.265)
Baseline Dummy: Year 6	-2.715	-2.339	-2.340	-2.260
	(0.313)	(0.340)	(0.341)	(0.348)
Baseline Dummy: Year 7	-3.886	-3.503	-3.498	-3.399
	(0.521)	(0.537)	(0.537)	(0.544)
Baseline Dummy: Year 8	-2.553	-2.187	-2.191	-2.077
	(0.524)	(0.542)	(0.542)	(0.545)
Dummy if Married	---	-0.180	-0.199	-0.170
		(0.124)	(0.127)	(0.128)
Dummy if White	---	-0.086	-0.088	-0.072
		(0.105)	(0.106)	(0.107)
Dummy if Age 30 to 34	---	-0.390	-0.396	-0.376
		(0.114)	(0.114)	(0.115)
Dummy if Age 35 to 39	---	-0.742	-0.757	-0.714
		(0.174)	(0.175)	(0.176)
Dummy if Some College	---	-0.054	-0.055	-0.035
		(0.103)	(0.103)	(0.103)
Dummy if College Degree	---	-0.137	-0.147	-0.137
		(0.102)	(0.103)	(0.103)

Table 4 (Continued)

Explanatory Variable	(1)	(2)	(3)	(4)
Dummy if Children Age 6 to 10	---	-0.169 (0.096)	-0.160 (0.096)	-0.134 (0.096)
Dummy if Children Age 11 to 18	---	-0.077 (0.121)	-0.072 (0.121)	-0.055 (0.121)
Dummy if Children Over Age 18	---	-0.368 (0.345)	-0.351 (0.343)	-0.264 (0.342)
Dummy if Not Employed	---	0.049 (0.109)	0.055 (0.109)	0.065 (0.109)
Dummy if Divorced	---	0.125 (0.183)	0.144 (0.183)	0.166 (0.184)
Dummy if Housing Expenditure-Constrained	---	---	-0.153 (0.114)	-0.133 (0.113)
Real Income	---	---	-0.00037 (0.00040)	-0.00034 (0.00040)
Real Net Worth	---	---	-0.00059 (0.00060)	0.00060 (0.00060)
County Unemployment Rate	---	---	---	-0.057 (0.021)
Log Likelihood	-1748.9	-1724.8	-1742.3	-1719.1
Estimated Percentage Shift in the Baseline Hazard due to Equity Constraint	-13.9	-12.7	-12.0	-10.5
Estimated Percentage Shift in the Baseline Hazard due to Nominal Loss	-41.6	-34.2	-33.5	-27.1
Estimated Percentage Shift in the Baseline Hazard due to Nominal Gain	-49.6	-39.2	-39.0	-37.2
p-Value for Test of Symmetry of Gains and Losses ($\delta = -\gamma$)	0.0001	0.0001	0.0001	0.006

Note: Prentice-Gloeckler-Meyer semiparametric hazard model estimates calculated on 6,461 household-year observations that comprise 3,112 residence spells, and 596 intrametropolitan own-to-own moves. The hazard is the probability of moving at time *t* conditional on not having moved before then. Standard errors are in parentheses. The *p*-values for the test of the null hypotheses of no equity constraint, no effect of nominal loss on mobility, and no effect of nominal gain on mobility (versus the alternatives outlined in the text) are shown in square brackets for the first three explanatory variables in the table, respectively. Income and net worth are in thousands of 1993 dollars. The county unemployment rate is measured in percentage points. The estimated percentage shift in the baseline hazard due to the equity constraint is calculated based on each specification's parameter estimate with respect to the interaction between the dummy if *LTV*>0.80 and the *LTV* (i.e., were in the constrained group), evaluated at the sample mean *LTV* for those in the constrained group of 0.905. The estimated

percentage shift in the baseline hazard due to the nominal loss is calculated based on each specification's parameter estimate with respect to the interaction between the dummy if nominal loss and the nominal loss, (the second explanatory variable in the table) evaluated at the sample mean nominal loss for those who experienced nominal losses of 0.054, or 5.4 percent loss. The estimated percentage shift in the baseline hazard due to the nominal gain is calculated based on each specification's parameter estimate with respect to the interaction between the dummy if nominal gain and the nominal gain, (the third explanatory variable in the table) evaluated at the sample mean nominal gain for those who experienced nominal gains of 0.135, or 13.5 percent gain. All specifications estimated with a full set of calendar year dummies.
Source: Author's calculation.

Table 5. Semiparametric Proportional Hazard Model Estimates of Intrametropolitan Own-to-Own Mobility, with Loan-to-ValueThreshold at 90 Percent

Explanatory Variable	(1)	(2)	(3)	(4)
Dummy if LTV>0.90	-0.203	-0.165	-0.139	-0.130
×LTV	(0.124)	(0.125)	(0.127)	(0.127)
	(0.051)	(0.093)	(0.138)	(0.154)
Dummy if Nominal Loss	-9.687	-7.616	-7.477	-5.785
×Nominal Loss	(2.458)	(2.463)	(2.463)	(2.478)
	(0.0001)	(0.001)	(0.001)	(0.010)
Dummy if Nominal Gain	-4.993	-3.665	-3.654	-3.433
×Nominal Gain	(0.528)	(0.569)	(0.571)	(0.573)
	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Baseline Dummy: Year 1	-3.748	-3.218	-3.201	-3.065
	(0.210)	(0.261)	(0.264)	(0.276)
Baseline Dummy: Year 2	-3.345	-2.864	-2.856	-2.749
	(0.214)	(0.259)	(0.262)	(0.272)
Baseline Dummy: Year 3	-2.951	-2.521	-2.510	-2.391
	(0.211)	(0.254)	(0.256)	(0.267)
Baseline Dummy: Year 4	-2.986	-2.534	-2.528	-2.426
	(0.234)	(0.275)	(0.276)	(0.286)
Baseline Dummy: Year 5	-2.675	-2.275	-2.267	-2.188
	(0.217)	(0.255)	(0.256)	(0.265)
Baseline Dummy: Year 6	-2.736	-2.357	-2.358	-2.277
	(0.313)	(0.340)	(0.341)	(0.348)
Baseline Dummy: Year 7	-3.907	-3.520	-3.515	-3.415
	(0.521)	(0.537)	(0.538)	(0.544)
Baseline Dummy: Year 8	-2.570	-2.199	-2.202	-2.088
	(0.524)	(0.542)	(0.542)	(0.545)
Dummy if Married	---	-0.189	-0.207	-0.177
		(0.124)	(0.126)	(0.128)
Dummy if White	---	-0.081	-0.086	-0.070
		(0.105)	(0.106)	(0.107)
Dummy if Age 30 to 34	---	-0.387	-0.395	-0.375
		(0.114)	(0.114)	(0.115)
Dummy if Age 35 to 39	---	-0.737	-0.754	-0.710
		(0.174)	(0.175)	(0.176)
Dummy if Some College	---	-0.057	-0.058	-0.037
		(0.103)	(0.103)	(0.103)
Dummy if College Degree	---	-0.136	-0.148	-0.137
		(0.102)	(0.103)	(0.103)

Table 5 (Continued)

Explanatory Variable	(1)	(2)	(3)	(4)
Dummy if Children Age 6 to 10	---	-0.167 (0.096)	-0.159 (0.096)	-0.132 (0.096)
Dummy if Children Age 11 to 18	---	-0.075 (0.121)	-0.070 (0.121)	-0.053 (0.121)
Dummy if Children Over Age 18	---	-0.371 (0.345)	-0.355 (0.344)	-0.267 (0.342)
Dummy if Not Employed	---	0.054 (0.109)	0.058 (0.109)	0.067 (0.109)
Dummy if Divorced	---	0.114 (0.182)	0.134 (0.183)	0.157 (0.184)
Dummy if Housing Expenditure-Constrained	---	---	-0.147 (0.114)	-0.128 (0.114)
Real Income	---	---	-0.00037 (0.00040)	-0.00034 (0.00040)
Real Net Worth	---	---	-0.00068 (0.00059)	0.00068 (0.00058)
County Unemployment Rate	---	---	---	-0.057 (0.020)
Log Likelihood	-1748.9	-1725.0	-1723.5	-1719.3
Estimated Percentage Shift in the Baseline Hazard due to Equity Constraint	-17.8	-14.8	-12.6	-11.7
Estimated Percentage Shift in the Baseline Hazard due to Nominal Loss	-40.7	-33.7	-33.2	-26.8
Estimated Percentage Shift in the Baseline Hazard due to Nominal Gain	-49.0	-39.0	-38.9	-37.1
p-Value for Test of Symmetry of Gains and Losses ($\delta = -\gamma$)	0.0001	0.0001	0.0001	0.007

Note: Prentice-Gloeckler-Meyer semiparametric hazard model estimates calculated on 6,461 household-year observations that comprise 3,112 residence spells, and 596 intrametropolitan own-to-own moves. The hazard is the probability of moving at time *t* conditional on not having moved before then. Standard errors are in parentheses. The *p*-values for the test of the null hypotheses of no equity constraint, no effect of nominal loss on mobility, and no effect of nominal gain on mobility (versus the alternatives outlined in the text) are shown in square brackets for the first three explanatory variables in the table, respectively. Income and net worth are in thousands

of 1993 dollars. The county unemployment rate is measured in percentage points. The estimated percentage shift in the baseline hazard due to the equity constraint is calculated based on each specification's parameter estimate with respect to the interaction between the dummy if *LTV*>0.90 and the *LTV* (i.e., were in the constrained group), evaluated at the sample mean *LTV* for those in the constrained group of 0.965. The estimated percentage shift in the baseline hazard due to the nominal loss is calculated based on each specification's parameter estimate with respect to the interaction between the dummy if nominal loss and the nominal loss, (the second explanatory variable in the table) evaluated at the sample mean nominal loss for those who experienced nominal losses of 0.054, or 5.4 percent loss. The estimated percentage shift in the baseline hazard due to the nominal gain is calculated based on each specification's parameter estimate with respect to the interaction between the dummy if nominal gain and the nominal gain, (the third explanatory variable in the table) evaluated at the sample mean nominal gain for those who experienced nominal gains of 0.135, or 13.5 percent gain. All specifications estimated with a full set of calendar year dummies.
Source: Author's caluculations.

Table 6. Semiparametric Proportional Hazard Model Estimates of Intrametropolitan Own-to-Rent Mobility, with Loan-to-Value Threshold at 80 Percent

Explanatory Variable	(1)	(2)	(3)	(4)
Dummy if LTV>0.80	-0.112	-0.050	-0.339	-0.329
×LTV	(0.179)	(0.181)	(0.191)	(0.191)
	(0.767)	(0.892)	(0.539)	(0.543)
Dummy if Nominal Loss	-9.765	-3.709	-3.149	-1.234
×Nominal Loss	(4.234)	(4.111)	(4.125)	(4.176)
	(0.011)	(0.183)	(0.222)	(0.384)
Dummy if Nominal Gain	-8.352	-5.318	-5.179	-4.814
×Nominal Gain	(1.069)	(1.190)	(1.204)	(1.229)
	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Baseline Dummy: Year 1	-5.103	-3.772	-3.544	-3.414
	(0.394)	(0.519)	(0.530)	(0.557)
Baseline Dummy: Year 2	-4.936	-3.751	-3.528	-3.423
	(0.406)	(0.516)	(0.525)	(0.549)
Baseline Dummy: Year 3	-4.599	-3.439	-3.247	-3.128
	(0.415)	(0.516)	(0.524)	(0.547)
Baseline Dummy: Year 4	-4.204	-3.068	-2.891	-2.798
	(0.422)	(0.521)	(0.529)	(0.550)
Baseline Dummy: Year 5	-4.490	-3.503	-3.349	-3.286
	(0.461)	(0.545)	(0.522)	(0.570)
Baseline Dummy: Year 6	-3.343	-2.346	-2.177	-2.100
	(0.492)	(0.569)	(0.574)	(0.592)
Baseline Dummy: Year 7	-3.422	-2.511	-2.366	-2.287
	(0.560)	(0.635)	(0.639)	(0.656)
Baseline Dummy: Year 8	-2.781	-1.826	-1.740	-1.623
	(0.770)	(0.811)	(0.823)	(0.836)
Dummy if Married	---	-0.629	-0.558	-0.527
		(0.210)	(0.214)	(0.217)
Dummy if White	---	-0.782	-0.718	-0.699
		(0.161)	(0.161)	(0.163)
Dummy if Age 30 to 34	---	-0.461	-0.389	-0.362
		(0.214)	(0.217)	(0.218)
Dummy if Age 35 to 39	---	-1.115	-1.003	-0.951
		(0.364)	(0.355)	(0.358)
Dummy if Some College	---	-0.404	-0.351	-0.343
		(0.178)	(0.179)	(0.179)
Dummy if College Degree	---	-1.290	-1.213	-1.218
		(0.247)	(0.248)	(0.248)

Table 6 (Continued)

Explanatory Variable	(1)	(2)	(3)	(4)
Dummy if Children Age 6 to 10	---	-0.185 (0.163)	-0.154 (0.164)	-0.183 (0.165)
Dummy if Children Age 11 to 18	---	-0.107 (0.202)	-0.068 (0.203)	-0.073 (0.203)
Dummy if Children Over Age 18	---	-0.529 (0.436)	-0.507 (0.441)	-0.646 (0.440)
Dummy if Not Employed	---	0.064 (0.199)	0.057 (0.200)	0.082 (0.201)
Dummy if Divorced	---	-0.075 (0.296)	-0.155 (0.298)	-0.145 (0.300)
Dummy if Housing Expenditure-Constrained	---	---	0.343 (0.184)	0.377 (0.185)
Real Income	---	---	-0.0012 (0.00069)	-0.0012 (0.00069)
Real Net Worth	---	---	-0.0076 (0.0018)	-0.0076 (0.0018)
County Unemployment Rate	---	---	---	-0.068 (0.036)
Log Likelihood	-789.4	-735.7	-725.4	-723.5
Estimated Percentage Shift in the Baseline Hazard due to Equity Constraint	-9.6	-4.4	-26.4	-25.7
Estimated Percentage Shift in the Baseline Hazard due to Nominal Loss	-41.0	-18.1	-15.6	-6.4
Estimated Percentage Shift in the Baseline Hazard due to Nominal Gain	-67.5	-51.1	-50.3	-15.3
p-Value for Test of Symmetry of Gains and Losses ($\delta = -\gamma$)	0.0001	0.053	0.076	0.207

Note: Prentice-Gloeckler-Meyer semiparametric hazard model estimates calculated on 6,461 household-year observations that comprise 3,112 residence spells, and 174 intrametropolitan own-to-rent moves. The hazard is the probability of moving at time *t* conditional on not having moved before then. Standard errors are in parentheses. The *p*-values for the test of the null hypotheses of no equity constraint, no effect of nominal loss on mobility, and no effect of nominal gain on mobility (versus the alternatives outlined in the text) are shown in square brackets for the first three explanatory variables in the table, respectively. Income and net worth are in thousands of 1993 dollars. The county unemployment rate is measured in percentage points. The estimated percentage shift in the baseline hazard due to the

equity constraint is calculated based on each specification's parameter estimate with respect to the interaction between the dummy if *LTV*>0.80 and the *LTV* (i.e., were in the constrained group), evaluated at the sample mean *LTV* for those in the constrained group of 0.905. The estimated percentage shift in the baseline hazard due to the nominal loss is calculated based on each specification's parameter estimate with respect to the interaction between the dummy if nominal loss and the nominal loss, (the second explanatory variable in the table) evaluated at the sample mean nominal loss for those who experienced nominal losses of 0.054, or 5.4 percent loss. The estimated percentage shift in the baseline hazard due to the nominal gain is calculated based on each specification's parameter estimate with respect to the interaction between the dummy if nominal gain and the nominal gain, (the third explanatory variable in the table) evaluated at the sample mean nominal gain for those who experienced nominal gains of 0.135, or 13.5 percent gain. All specifications estimated with a full set of calendar year dummies.
Source: Author's calculations.

Table 7. Semiparametric Proportional Hazard Model Estimates of Intermetropolitan Mobility, with Loan-to-Value Threshold at 80 Percent

Explanatory Variable	(1)	(2)	(3)	(4)
Dummy if LTV>0.80	0.113	0.131	0.152	0.167
×LTV	(0.209)	(0.212)	(0.221)	(0.222)
	(0.588)	(0.538)	(0.492)	(0.451)
Dummy if Nominal Loss	-6.878	-0.273	-0.267	2.208
×Nominal Loss	(4.491)	(4.200)	(4.231)	(4.236)
	(0.062)	(0.474)	(0.475)	(0.602)
Dummy if Nominal Gain	-6.768	-3.753	-3.682	-3.200
×Nominal Gain	(1.012)	(1.176)	(1.168)	(1.187)
	(0.0001)	(0.001)	(0.001)	(0.007)
Baseline Dummy: Year 1	-5.275	-4.410	-4.394	-4.145
	(0.407)	(0.498)	(0.501)	(0.528)
Baseline Dummy: Year 2	-4.573	-3.868	-3.839	-3.624
	(0.412)	(0.489)	(0.491)	(0.513)
Baseline Dummy: Year 3	-4.261	-3.619	-3.600	-3.387
	(0.411)	(0.481)	(0.482)	(0.504)
Baseline Dummy: Year 4	-4.202	-3.565	-3.553	-3.375
	(0.428)	(0.499)	(0.499)	(0.516)
Baseline Dummy: Year 5	-4.563	-3.994	-3.993	-3.847
	(0.505)	(0.554)	(0.554)	(0.567)
Baseline Dummy: Year 6	-4.478	-3.927	-3.909	-3.762
	(0.784)	(0.815)	(0.815)	(0.823)
Baseline Dummy: Year 7	-4.076	-3.547	-3.529	-3.338
	(0.745)	(0.775)	(0.775)	(0.785)
Baseline Dummy: Year 8	-19.679	-20.220	-20.205	-20.079
	(2109)	(3509)	(3509)	(3503)
Dummy if Married	---	-0.463	-0.515	-0.447
		(0.231)	(0.235)	(0.241)
Dummy if White	---	-0.184	-0.213	-0.172
		(0.218)	(0.221)	(0.225)
Dummy if Age 30 to 34	---	-0.451	-0.478	-0.434
		(0.234)	(0.236)	(0.238)
Dummy if Age 35 to 39	---	-1.104	-1.137	-1.043
		(0.395)	(0.396)	(0.400)
Dummy if Some College	---	-0.364	-0.386	-0.338
		(0.263)	(0.264)	(0.266)
Dummy if College	---	0.569	0.527	0.543
Degree		(0.204)	(0.207)	(0.208)

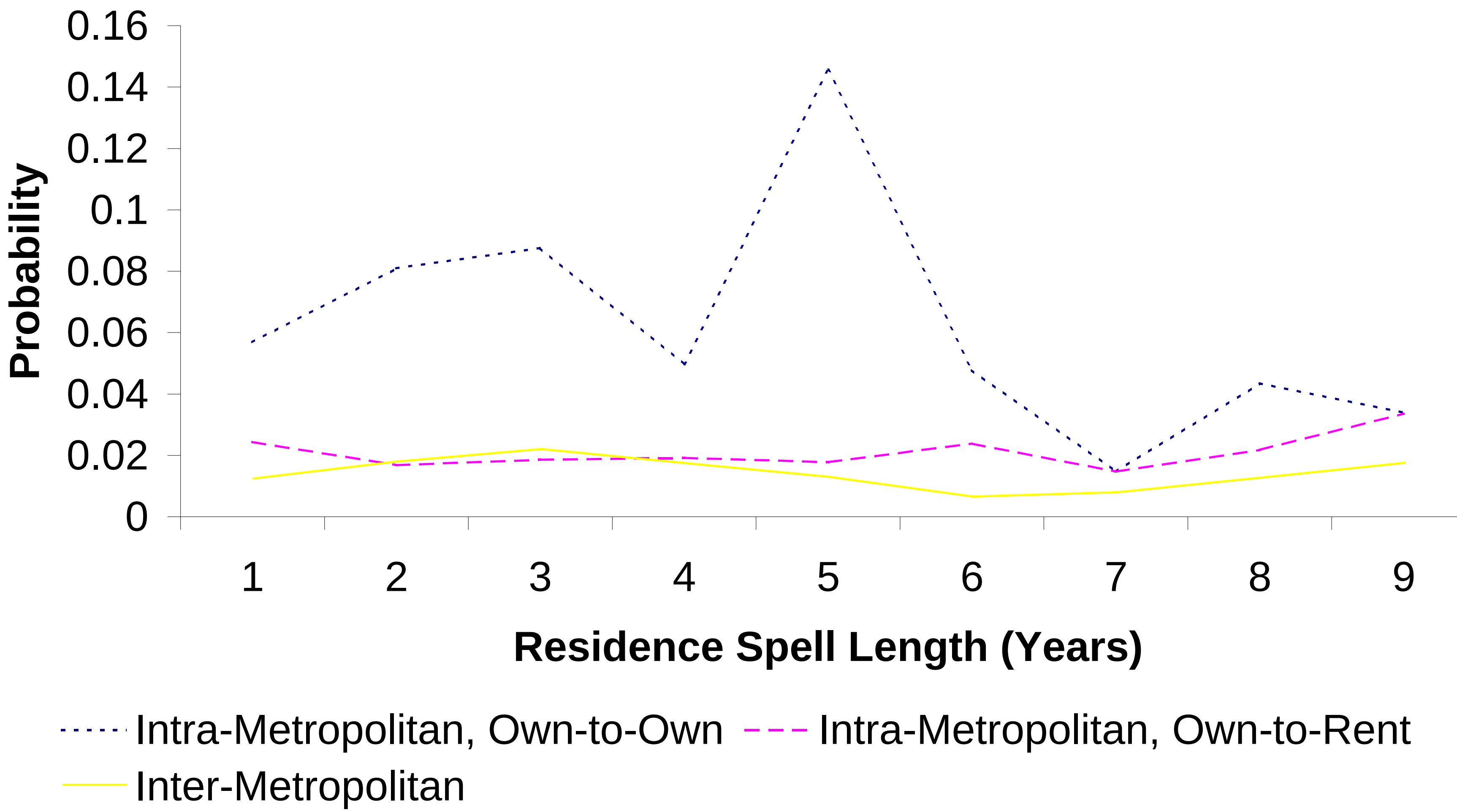
Table 7 (Continued)

Explanatory Variable	(1)	(2)	(3)	(4)
Dummy if Children Age 5 and Under	---	-0.152 (0.198)	-0.150 (0.198)	-0.138 (0.199)
Dummy if Children Age 6 to 10	---	0.032 (0.217)	0.039 (0.217)	0.088 (0.218)
Dummy if Children Age 11 to 18	---	-0.098 (0.282)	-0.098 (0.283)	-0.064 (0.283)
Dummy if Children Over Age 18	---	-17.319 (2189)	-17.285 (2190)	-17.077 (2208)
Dummy if Not Employed	---	0.417 (0.213)	0.413 (0.214)	0.425 (0.215)
Dummy if Divorced	---	-0.497 (0.432)	-0.494 (0.433)	-0.430 (0.435)
Dummy if Housing Expenditure-Constrained	---	---	-0.041 (0.235)	-0.005 (0.237)
Real Income	---	---	-0.00086 (0.00055)	0.00088 (0.00054)
Real Net Worth	---	---	-0.00055 (0.0012)	0.00057 (0.0012)
County Unemployment Rate	---	---	---	-0.094 (0.044)
Log Likelihood	-600.6	-575.5	-574.1	-571.5
Estimated Percentage Shift in the Baseline Hazard due to Equity Constraint	10.8	12.6	-12.9	16.3
Estimated Percentage Shift in the Baseline Hazard due to Nominal Loss	-31.0	-1.5	-1.4	12.7
Estimated Percentage Shift in the Baseline Hazard due to Nominal Gain	-59.8	-39.7	-39.2	-35.0
p-Value for Test of Symmetry of Gains and Losses ($\delta = -\gamma$)	0.005	0.394	0.405	0.836

Note: Prentice-Gloeckler-Meyer semiparametric hazard model estimates calculated on 6,461 household-year observations that comprise 3,112 residence spells, and 125 intermetropolitan moves. The hazard is the probability of moving at time *t* conditional on not having moved before then. Standard errors are in parentheses. The *p*-values for the test of the null hypotheses of no equity constraint, no effect of

nominal loss on mobility, and no effect of nominal gain on mobility (versus the alternatives outlined in the text) are shown in square brackets for the first three explanatory variables in the table, respectively. Income and net worth are in thousands of 1993 dollars. The county unemployment rate is measured in percentage points. The estimated percentage shift in the baseline hazard due to the equity constraint is calculated based on each specification's parameter estimate with respect to the interaction between the dummy if *LTV*>0.80 and the *LTV* (i.e., were in the constrained group), evaluated at the sample mean *LTV* for those in the constrained group of 0.905. The estimated percentage shift in the baseline hazard due to the nominal loss is calculated based on each specification's parameter estimate with respect to the interaction between the dummy if nominal loss and the nominal loss, (the second explanatory variable in the table) evaluated at the sample mean nominal loss for those who experienced nominal losses of 0.054, or 5.4 percent loss. The estimated percentage shift in the baseline hazard due to the nominal gain is calculated based on each specification's parameter estimate with respect to the interaction between the dummy if nominal gain and the nominal gain, (the third explanatory variable in the table) evaluated at the sample mean nominal gain for those who experienced nominal gains of 0.135, or 13.5 percent gain. All specifications estimated with a full set of calendar year dummies.
Source: Author's calculations.

Figure 1. Kaplan-Meier Empirical Hazards by Type of Move



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